J. Forest Economics 9, 97–117 (2003)

http://www.urbanfischer.de/journals/jfe

Journal of FOREST ECONOMICS

Market structure in U. S. southern pine roundwood

Matthew F. Bingham[a], Jeffrey P. Prestemon[b], Douglas J. MacNair[a], and Robert C. Abt[c]

[a] Triangle Economic Research, 2775 Meridian Parkway, Durham, North Carolina, 27713, USA
[b] Forestry Sciences Laboratory, Southern Research Station, USDA Forest Service, P. O. Box 12254, Research Triangle Park, North Carolina, 27709, USA
[c] Department of Forestry, North Carolina State University, Box 8008, Raleigh, North Carolina, 27695-8008, USA

Abstract

Time series of commodity prices from multiple locations can behave as if responding to forces of spatial arbitrage, even while such prices may instead be responding similarly to common factors aside from spatial arbitrage. Hence, while the Law of One Price may hold as a statistical concept, its acceptance is not sufficient to conclude market integration. We tested the factors hypothesized as linked to integration of forest products markets by applying a combination of bivariate and multivariate techniques. Bivariate cointegration tests were conducted for price pairs among 21 price regions and were done for both delivered southern pine sawlogs and delivered southern pine pulpwood logs. Multivariate meta-analytic regressions of cointegration test results on hypothesized explanatory factors were run for pulpwood and sawlog markets separately. Cointegration test results offer limited support for the Law of One Price in the South for both products. Results of the meta-analytic regressions show that a proxy for the cost of product transfer between regions is statistically significant and negatively related to the probability that two local market prices are cointegrated for only sawlogs. For pulpwood, the proxy was not significant. The results of the bivariate cointegration tests and the multivariate meta-analyses were used to delineate apparently spatially segmented sub-markets for both products. The maps show overlapping geographical segments, resulting from both spatial arbitrage and possible output dominance for certain firms in those sub-markets. The southern pine sawlog market can be divided into four or five sub-markets, distributed north to south and east to west. The southern pine pulpwood log market can be drawn into three, largely separate sub-markets: a coastal zone that stretches from Texas to Virginia, and two distinct interior zones.

Key words: Law of One Price, log markets, meta-analysis, cointegration, pine

JEL classification: Q210, D400

Introduction

Recognition of regional resource characteristics and associated markets is a defining component of forest sector market models. These models, used to project future conditions of the forest and forest product markets, typically are either re-

Correspondence to: Matthew F. Bingham, Triangle Economic Research, 2775 Meridian Parkway, Durham, North Carolina, 27713, USA. E-mail: mbingham@ter.com

1104-6899/03/09/02-097 $ 15.00/0

gion-specific, modeling the effects of broad market variables on subcomponents within the region, or spatially linked regional models. Region-specific models include the Subregional Timber Supply Model, SRTS (Abt et al. 2000), which is used to project timber supply and demand conditions within the U. S. South. Spatially linked models include the Timber Assessment Market Model, TAMM (Adams and Haynes 1980), which in updated form has been used in successive Resource Planning Act-mandated projections of U. S. national supply and demand conditions for specific regions in the U. S. Another is the Global Trade Model (Dykstra and Kallio 1987), used to model world forest product markets.

A common assumption of forest sector projection models, aside from perfect competition in timber markets, is that all points of production and consumption of the same product are parts of the same market within a region modeled, implying both efficient intra-regional market shock transmission as well as the existence of a single price – i. e., the applicability of the Law of One Price (LOP). However, if markets do not transmit shocks, then prices in one part of the producing market do not respond to price fluctuations in other parts, failing the conditions needed for market integration (Ravallion 1986). In the absence of price shock transmission, sub-optimal marketing and production decisions and wealth transfers can result.

Empirical support for integration of forest product markets is limited. In the lowest stages of forest product output, the results are negative. Nagubadi et al. (2001) have shown that the single market assumption, which they measured with cointegration of multiple series of hardwood stumpage prices, did not apply to southern U. S. hardwood stumpage markets. Prestemon and Holmes (2000) conducted market shock price imprint tests that suggested southern pine stumpage markets are not integrated; bivariate cointegration tests of prices did not support the Southwide applicability of the LOP. Studies of price behavior in markets of forest products in higher stages of production have been used to evaluate both the LOP and, it is sometimes claimed, market integration. Most of these studies supported the LOP. Examples include the causality tests of Uri and Boyd (1990) and the cointegration testing of many others: Jung and Doroodian (1994) and Murray and Wear's (1998) examinations of U. S. softwood lumber markets, Buongiorno and Uusivuori's (1992) evaluation of U. S. pulp and paper exports, Alavalapati et al.'s (1997) study of Canada's pulpwood market, Hänninen et al.'s (1997) analysis of newsprint price behavior in northern Europe and Canada, and Toppinen and Toivonen's (1998) examination of the Finnish pulpwood market.

Cointegration testing has been applied in many of the aforementioned studies and in other commodity markets (e. g., Ardeni 1989; Baffes 1991) as the means for evaluating whether the LOP holds, but cointegration test results often provide limited information about the mechanisms of market integration. McNew and Fackler (1997), who distinguish between market efficiency (as embodied by the LOP) and market integration (as embodied by local market shock transmission), consistent with Ravallion (1986), suggested that cointegration can occur even in the absence of direct spatial arbitrage in markets for products within one stage of processing – for example, through market integration in spatially distributed higher stages of production. Consistent with that idea, Prestemon and Holmes (2000) suggested that it was the integration of markets in higher stages of wood processing that helped explain the results of cointegration in pulpwood stumpage market prices, where cointegration did not seem to be related to distance.

Nevertheless, the possibilities remain that the LOP as measured by cointegration testing can hold even when markets are not integrated, and integration can occur when price behavior tests such as cointegration indicate that the LOP does not hold (Söderlind and Vredin 1996; McNew and Fackler 1997; Phillips 1998). These facts make interpretation of cointegration test results problematic. McNew and Fackler (1997) state and show how integrated markets can have price series that fail cointegration tests for a variety of other reasons. For example, cointegration tests can provide the wrong answer if the analyst does not account for needed trends in relationships between price series or inserts trend variables when trends actually do not exist. Similar cointegration testing mistakes can be made when considering log-transforming data, accounting for inflation, addressing subtle product quality differences, and accounting for possible market structural changes. Finally, the presence of transaction costs for product movements, often not observed by the analyst but which can vary independently from product prices over time, can lead to incorrect conclusions about market integration (Barrett, 1996). Incorrectly accounting for the effects of variable transaction or product transfer costs can lead to incorrect inferences (Baulch 1997; Murray and Wear 1998).

Our central goal is to better understand southern pine log market integration. Mitigating some of the possible pitfalls in LOP and market integration testing in our examination of the structure of southern pine log markets, we combine both price cointegration tests and an approach to evaluate whether incomplete market integration can explain these results. To accomplish this, we adapt the techniques of Goodwin and Schroeder (1991), who bridged the gap between the LOP and market integration by explaining the results of cointegration tests using auxiliary regressions. In empirical testing, if cointegration of prices can be tied statistically to variables directly related to product movement, then cointegration of prices may in fact be consistent with market integration because product movement or the threat of profitable product movement is required for shock transmission in the form of price adjustments (Takayama and Judge 1964). If other variables – perhaps variables linked to output markets or common policy shocks – are statistically related to cointegration, then the results of Prestemon and Holmes (2000) and the critiques of Söderlind and Vredin (1996), McNew and Fackler (1997), and Phillips (1998) would be more strongly supported. The scope of our analysis is Southwide, using information from price series generated by the log sale activities at mills in geographically small southern pine sawlog and pulpwood regions. Results show that the LOP seems to fail and that Southwide southern pine log markets are not well integrated. The lack of integration may be caused by the high relative expense of spatial arbitrage. Our findings are synthesized to delineate regional sub-markets, within which both the LOP and possibly "sub-market integration" may be comfortably assumed to exist.

Methods

Our choice of cointegration test is the two-step estimator of Engle and Granger (1987), testing for stationarity of the residuals of linear regressions of one price on another, which, we explain later, are the most appropriate tests for this analysis. Because alternative specifications of cointegrating relations often return divergent results for the same price series, we apply a heuristic approach to identifying the appropriate specification of the cointegrating regression. The bivariate price regressions of the OLS two-step approach use alternative specifications of the cointegrating relation, including testing for whether the cointegrating relation should include a policy shift variable, with or without inflation removed, and with or without prices transformed by the natural logarithm. Each cointegrating relation specification for each price pair returns a p-value representing the probability that the two series are cointegrated in the manner supposed by the specified relation. The set of p-values generated from tests of cointegration for each specification of the cointegrating relation are then used in meta-analytic regressions of cointegration test results on hypothesized explanatory variables.

Meta-analysis, a statistical technique for synthesizing diverse results from different studies of the same response variable (Hedges and Olkin 1985), is the means by which we evaluate whether arbitrage or other variables are linked to the probability that two variables possess a linear cointegrating relation. Many applications of meta-analysis have involved the health sciences, but the approach also is applied in economics (e. g., Smith and Kaoru 1990, Goodwin and Schroeder 1991, Espey 1998, Stanley 1998). The idea is straightforward. An analyst may have findings from several studies about the estimated size of a particular parameter or variable, a "response variable," but recognize that the findings were generated by studies with a diverse set of testing conditions or modeling assumptions. Meta-analytic techniques can tease out the individual effects of these alternative test conditions or assumptions on the estimate of the response variable. A quantitative approach to identifying their effects is by estimating an equation, $Y_i = f(S_i, M_i) + \varepsilon_i$, where Y_i is the response variable of interest for study i, S_i is a vector of the characteristics of sample i from which the statistic was generated, and M_i are the modeling assumptions of study i. The exact functional form depends on the specific application.

Cointegration

Economic theory often suggests that underlying forces unify a pair of economic variables (Engle and Granger 1987). The influence of arbitrage on prices of similar commodities in spatially separated markets is cited as the primary mechanism for price unification (Ardeni 1989; Goodwin and Schroeder 1991). Here, the competitive actions of profit-seeking individuals link regional markets. The purchase of commodities in lower-priced markets combined with their subsequent transportation to and resale in higher-priced markets results in profit for commodity arbitrageurs (Samuelson 1952; Takayama and Judge 1964). Competition among arbitrageurs quickly eliminates excess profits arising from such activities. Thus, economic theory predicts an equilibrium relationship such that price differentials between two markets are no greater than the sum of transportation costs, transaction costs, and some nonnegative profit. The existence of intermediate markets implies reduced transportation costs and subsequently finer price differentials (Mulligan and Fik 1989).

If arbitrage is the mechanism by which prices for identical products in two spatially diverse regions are unified and market shocks in one part of the region necessarily are transmitted to other parts of the region, then we can say that the LOP applies to the two regions and that the two are integrated into one market. When arbitrage keeps these prices together, price differences just offset transfer (transportation plus sale transaction) costs. If the price in region H is P^H and in region L is P^L, then their difference in period t is $P_t^H - P_t^L = T_t^{HL}$. Here, efficiently linked regions mean that the price difference, T_t^{HL}, is identical to transfer costs. Thus, some of region L's excess supply at P^L will be shipped to region H and sold at price P^H. Product movement occurs until the price difference between the two regions is no more than the cost of transfer (T^{HL}).

Cointegration provides a means to test for price linkages. If two prices are stationary, then one price can always be expressed as a linear combination of the other, and the LOP must be evaluated through means such as a vector autoregression in differences. If two prices are nonstationary, of the same order of integration, and their linear combination is stationary, then these series possess a cointegrating relation. Consider a vector of $I(1)$ variables x_t. If there exists a vector β that, when multiplied times x_t, creates another series μ_t that is $I(0)$, then the elements of x_t are cointegrated with a cointegrating relation β:

$$x_t' \beta = \mu_t \qquad [1]$$

A simple cointegrated bivariate system can be expressed as

$$\begin{aligned} y_t &= \lambda x_t + u_{yt} \\ x_t &= x_{t-1} + u_{xt} \end{aligned} \qquad [2]$$

Here, u_{yt} and u_{xt} are uncorrelated white noise and x_t is a random walk. From here on, we confine our discussion to the bivariate case of cointegration. Following Engle and Granger (1987), testing for cointegration in the bivariate case involves a set of univariate tests on each variable and then two OLS regressions. The univariate test is to evaluate whether two series are integrated of the same order. A common test is the augmented Dickey-Fuller test (Dickey and Fuller 1979, Said and Dickey 1984). In this research, we employed the weighted symmetric test for stationarity (Pantula et al. 1994) because it is more powerful than the unit root test developed by Dickey and Fuller (1979). Next, Engle and Granger (1987) outline two OLS regressions (hence, their "two-step" title to the approach):

- Step 1: Estimate a hypothesized cointegrating relation by regressing one variable on the other, under some assumed form of the relation.
- Step 2: Retain the residuals from the first regression and test them for nonstationarity.

Critical values for nonstationarity exist for augmented Dickey-Fuller regressions from the response surface of critical values derived by MacKinnon (1991).

When testing for cointegration using the OLS method, the resultant test statistic typically depends upon which price series is chosen as the independent variable and which series is the dependent variable. Johansen and Juselius (1990) and Johansen (1991) suggest avoiding this difficulty by employing maximum likelihood in an error correction model (MLECM), to force price relationship symmetry. Although the MLECM method is generally considered the "state of the art" technique for evaluating cointegration, it is not without limitations. In particular for this research, the MLECM approach would limit rather than strengthen our analysis.

The primary reason for not employing MLECM in this analysis is that the well-known asymmetry of the results obtained from the alternate selection of which variable is in the "dependent" position in two-step tests is desirable. Specification asymmetry means that employing the two-step test in place of the FIML test allows us to test for the influence of factors unique to each market on cointegration test results. As the MLECM test forces the same result for each specification, employing this test would weaken the meta-analysis in two important ways. First, the MLECM specification does not recognize "home" and "away" markets. Thus, a meta-analysis employing MLECM cointegration results could not evaluate the influence of volume and concentration in "home" and "away" markets. Further, allowing for the possibility that one series appears differently integrated with another provides us with an opportunity to determine why OLS cointegration tests might give divergent results depending on which variable is in the dependent position and which one is in the independent position in the regression. Second, the assumption of price relationship symmetry means that employing the MLECM test halves the number of cointegration results available for analysis.

We also considered the additional power that is supposed to be provided by the MLECM test. Recent evidence, based on simulations, suggests that the MLECM test provides only marginally greater testing power and accuracy of cointegrating parameters over the OLS approach (Gonzalo 1994). Further, the consequences of under-parameterization for MLECM may be as deleterious as they are for underparameterizing the nonstationarity test of the second step of the two-step approach. Arbitrarily truncating the number of lagged difference terms in the vector autoregression component of the MLECM – whose inclusion helps to reduce bias – may result in incorrect inference more often than in the OLS approach. Additionally, the higher number of parameters sometimes included in the MLECM approach can weaken statistical inference, raising the probability of type-I errors regarding the null of no cointegration.

We chose cointegrating relation specifications in order to evaluate whether the results of cointegration were sensitive to policies, inflation, and data transformation. The policies that we tested were associated with the reductions in timber harvest on federal lands, which took effect in 1988 and continue to the present (Murray and Wear 1998). This policy affected tests of cointegration in the Murray and Wear (1998) study and was included in analyses by Prestemon and Holmes (2000) and Guan and Munn (2000). Alternative specifications of the OLS cointegration equations therefore include and exclude a dummy variable that is zero before 1988, one after that.

Schnute (1987) warns that imposing a filtering process such as deflating prices can result in spurious patterns and spuriously significant relationships among variables. Believing that not removing inflation from prices carries its own risks of spurious findings of cointegration, alternative specifications of the cointegration equation alternately deflate and do not deflate the price variables in the bivariate cointegration tests using the producer price index (PPI) for all commodities.

Finally, analyses in economics often have shown that prices are distributed lognormally, rather than normally in untransformed form. Heteroscedasticity, especially arising from inflation, can be limited by transforming prices. In our analysis, alternative specifications of cointegrating relations express variables in logged and unlogged forms. Logarithmic transformation also removes the wedge effect of inflation on nominal prices, which would tend to drive the prices progressively further apart in the long run.

In total, then, there are 8 different specifications of cointegrating relations for prices from 21 local markets. Each cointegrating relation specification includes an intercept. In the augmented Dickey-Fuller test equations of the residual series deriving from the estimating of each cointegrating relation, the number of lagged difference terms included was that which minimized the Akaike Information Criterion, a model selection procedure recommended by Hall (1994). Because regressions are done with prices in both the "dependent" and the "independent" sides of OLS regressions, we have $21 \times 20 = 420$ observations in each meta-regression in the meta-analysis stage of our study.

Meta-analysis of cointegration test statistics

Meta-analytic regressions are used in our research to find evidence supporting one or more cointegrating relation specifications and to identify the reasons why two prices may or may not be cointegrated. If the explanatory variables in the meta-regressions, which in our case are static economic variables describing the markets whose prices are being compared, affect the probability of cointegration in the manner hypothesized a priori, then the original cointegrating relation applied in the Engle-Granger cointegration tests would appear favored. In a similar vein, if a particular cointegrating relation specification yields statistics that are unrelated or related to the static economic variables in ways not expected from theory, it is more likely that the p-values did not arise from the "correct" cointegrating relation specification. If p-values from cointegration tests on all of the most plausible cointegrating relations specified fail to support cointegration, then the approach would have largely failed to accept or reject concepts of the LOP or market integration. Thus, this approach provides guidance in selecting the appropriate cointegrating relation, and the best-fitting meta-analytic regression would appear to be the one that best describes the structure behind interregional market-price dynamics.

Identification of the set of plausible cointegrating relations applicable to southern pine log prices in the South was based on the concept of market integration, economic theory, and empirical evidence of other research. Market shock transmission may be limited by the cost of spatial arbitrage (Takayama and Judge 1964, Mulligan and Fik 1989). Not possessing exact information on the cost of transfer of product between two production/consumption points (regions) and given the concept of spatial arbitrage, distance between two regions should be negatively related to the probability of price cointegration. Faminow and Benson (1990) also suggest the potential existence of basing-point pricing systems, whereby prices from two or more regions are linked because of collusive behavior of producers. Hence, with high concentration of production in the home region, it would be more probable in cases of incomplete support for the LOP through cointegration that a comparison region's prices maintain a stable relationship with that home region. Lang and Rosa (1981) and Buccola (1985) have hypothesized that localities with greater economic activity (i. e., more transactions) have more efficient price-generating mechanisms because information about prices is more frequently observed and readily available. But according to Tomek (1980), regions with low production volumes may exhibit price swings not in line with other regions. Such low volume regions might also exhibit large, nonstationary shocks because of the start-up or shutdown of individual production facilities, meaning a permanent departure from past relationships between prices from other regions. Hence, the volume of production in the home region should be positively related with the probability of finding a significant cointegrating relationship with other, spatially separated regions. But as Goodwin and Schroeder (1991) found for regional cattle markets, the volume relationship is not clear. We would, therefore, have no a priori expectation of the direction of influence of market concentration on the probability of cointegration.

The basic form of the meta-regression estimated here is:

$$COINT_{yx} = c_0 + c_1\, DIST_{yx} + c_2 VOL_y + c_3 VOL_x + c_4 CONC_y + c_5 CONC_x + \upsilon_{yx} \quad [3]$$

Where

$COINT_{yx}$ = the p-values arising from the augmented Dickey-Fuller tests of stationarity of residuals of the hypothesized cointegration equation, arising from the OLS regression of y_t on x_t

$DIST_{yx}$ = the distance, in miles, between market centers of y_t and x_t

VOL_y = average volume from region y

VOL_x = average volume from region x

$CONC_y$ = number of mills in region y

$CONC_x$ = number of mills in region x

The meta-analytic regressions are estimated using minimum chi-squared econometric techniques. Minimum chi-squared regression is a form of weighted ordinary least squares, which recognizes that the dependent variable is a probability.

Data

The source of southern pine sawlog and pulpwood log price data were the nominal quarterly price series of delivered southern pine sawlogs and pulpwood logs corresponding to 21 of the price regions (with long enough series to permit analysis) of Timber Mart-South (TMS) (Norris Foundation 1977–1999). These data spanned the period from the first quarter of 1977 to the fourth quarter of 1998, resulting in 88 observations per series. TMS presently reports price data for 2 price regions per state, and our 21 series corresponded to the price regions of 11 states (excluding 1 price region in Arkansas and both in Oklahoma and Kentucky). But TMS's price regions and the temporal frequency of reports have changed over time. Before 1992, there were sometimes three regions per state. To correct this spatial inconsistency, we employed the conversion procedure recommended by Prestemon and Pye (2000). Before 1988, prices were reported monthly for each price region. To address this temporal inconsistency in our data, we adopted the recommendations of Haight and Holmes (1991), in our case taking middle-month sample by quarter from monthly series for 1977–1987.

For the meta-analysis, distance between markets was measured from the geographic center of each TMS region using PCMiler©, a software program commonly used by trucking companies for routing trips. Timber output and mill count data, used to calculate an industrial concentration index for each price region, were obtained from a variety of sources. Data for Texas and Louisiana on timber removals (sawtimber and pulpwood), our chosen index of market activity, and mills by region were obtained by special request from the Forest Inventory and Analysis unit of the United States Department of Agriculture-Forest Service, Asheville, North Carolina. The same variables were obtained for the remaining states in our study from reports by Howell and Levins (1998), Johnson et al. (1997a–f), Stratton and Wright (1998) and Stratton et al. (1998).

Empirical Results

Unit root tests on southern pine pulpwood and sawlog price series from the 21 TMS regions analyzed using techniques of Pantula et al. (1994) verified that each price series for each product is nonstationary. A table of these results is available from the authors. Nonstationary price behavior could not be rejected for any series, regardless of whether series were transformed by the natural logarithm or deflated, and regardless of whether a structural change variable corresponding with the Pacific Northwest harvesting restrictions was included in the regression. Confirmation of nonstationary price behavior permitted further analysis of price links using cointegration techniques.

Results of the two-step OLS cointegration tests are reported in Table 1 (delivered southern pine sawlogs) and Table 2 (delivered southern pine pulpwood). Results are presented for only the specification in which nominal prices were transformed by the natural logarithm and no structural change variable associated with the Pacific Northwest set-asides was included. Values in the tables are the p-values of cointegration. The smaller the value, the greater the likelihood that the residual series in the second stage of the two-step procedure (the augmented Dickey-Fuller tests of residuals) was stationary.

Both of these tables support the notion that the LOP does not hold from one end of the southern pine log market to the other within either product category. In Table 1, sawlog prices are cointegrated at 5 % significance in about one-third of price pairs. At 10 % significance, the proportion increases to about 42 %. For pulpwood, the proportion of price pairs with statistically significant cointegration at 5 % significance was about 18 %; with significance set at the weaker threshold of 10 % significance, only in 28 % of the price pairs could the null of no cointegration be rejected. Clearly, if cointegration tests were used as the criteria for determining whether the LOP holds in the South, then our results would suggest that it does not.

Meta-analysis of cointegration test result

Delivered southern pine sawlogs: The results from minimum chi-squared regressions of sawlog price cointegration p-values are shown in Table 3. The results from the meta-analysis show that economic considerations are statistically related to whether or not markets are cointegrated. In most cases, the parameter estimates associated with hypothesized variables were statistically different from zero when the signs were in the directions expected. Distance had a negative relationship with the probability of cointegration estimate in all specifications – the greater the

Table 1. Results of bivariate cointegration tests for delivered southern pine sawlogs, for the cointegrating relation specified as containing an intercept but no policy shock. Prices were undeflated prices and were transformed by the natural logarithm. Postal abbreviation codes were used to identify states and numbers (1, 2) to identify Timber Mart-South price regions within states.

X/Y	AL1	AL2	AR1	FL1	FL2	GA1	GA2	LA1	LA2	MS1	MS2	NC1	NC2	SC1	SC2	TN1	TN2	TX1	TX2	VA1	VA2
AL1	na	0.72	0.82	0.43	0.08	0.56	0.14	0.80	0.37	0.34	0.23	0.01	0.00	0.02	0.03	0.00	0.00	0.85	0.76	0.01	0.00
AL2	0.72	na	0.42	0.08	0.01	0.54	0.02	0.60	0.30	0.59	0.30	0.00	0.08	0.03	0.00	0.05	0.00	0.29	0.33	0.83	0.63
AR1	0.70	0.24	na	0.27	0.07	0.21	0.03	0.03	0.00	0.04	0.00	0.05	0.94	0.04	0.12	0.00	0.01	0.10	0.09	0.08	0.67
FL1	0.10	0.05	0.31	na	0.11	0.54	0.25	0.85	0.45	0.42	0.19	0.00	0.39	0.03	0.00	0.00	0.46	0.37	0.25	0.64	0.12
FL2	0.30	0.01	0.18	0.05	na	0.49	0.00	0.64	0.26	0.34	0.20	0.01	0.07	0.02	0.01	0.30	0.56	0.29	0.15	0.55	0.37
GA1	0.53	0.25	0.23	0.41	0.24	na	0.16	0.61	0.56	0.58	0.24	0.02	0.45	0.00	0.02	0.00	0.25	0.65	0.59	0.97	0.57
GA2	0.08	0.01	0.04	0.04	0.00	0.20	na	0.77	0.63	0.43	0.08	0.03	0.10	0.01	0.00	0.00	0.00	0.42	0.44	0.17	0.62
LA1	0.16	0.13	0.01	0.20	0.06	0.32	0.42	na	0.05	0.01	0.00	0.02	0.23	0.02	0.09	0.00	0.47	0.00	0.01	0.37	0.61
LA2	0.14	0.08	0.00	0.26	0.06	0.32	0.33	0.24	na	0.01	0.00	0.01	0.35	0.03	0.01	0.43	0.51	0.01	0.02	0.45	0.69
MS1	0.11	0.10	0.01	0.16	0.04	0.35	0.18	0.02	0.01	na	0.02	0.03	0.04	0.02	0.09	0.50	0.49	0.21	0.01	0.05	0.53
MS2	0.21	0.08	0.01	0.21	0.11	0.31	0.10	0.04	0.00	0.10	na	0.04	0.72	0.03	0.10	0.54	0.00	0.01	0.00	0.62	0.65
NC1	0.85	0.86	0.87	0.79	0.90	0.81	0.83	0.87	0.94	0.88	0.77	na	0.82	0.20	0.67	0.94	0.05	0.85	0.93	0.99	0.84
NC2	0.07	0.21	0.75	0.41	0.10	0.49	0.50	0.70	0.72	0.73	0.26	0.03	na	0.19	0.17	0.02	0.02	0.84	0.88	0.04	0.02
SC1	0.10	0.13	0.27	0.20	0.07	0.19	0.07	0.63	0.33	0.29	0.13	0.01	0.66	na	0.02	0.01	0.00	0.84	0.78	0.04	0.27
SC2	0.13	0.01	0.50	0.02	0.00	0.59	0.00	0.76	0.95	0.69	0.37	0.01	0.73	0.03	na	0.00	0.00	0.61	0.79	0.81	0.76
TN1	0.12	0.01	0.17	0.03	0.00	0.14	0.06	0.65	0.67	0.91	0.03	0.00	0.11	0.06	0.01	na	0.31	0.75	0.24	0.90	0.56
TN2	0.18	0.11	0.34	0.81	0.04	0.34	0.42	0.94	0.93	0.93	0.10	0.00	0.19	0.03	0.01	0.03	na	0.94	0.90	0.25	0.40
TX1	0.16	0.06	0.03	0.13	0.03	0.32	0.04	0.01	0.01	0.20	0.01	0.04	0.06	0.03	0.01	0.00	0.52	na	0.05	0.89	0.62
TX2	0.17	0.06	0.03	0.07	0.01	0.34	0.12	0.01	0.01	0.01	0.01	0.04	0.38	0.05	0.01	0.00	0.45	0.04	na	0.91	0.63
VA1	0.02	0.15	0.54	0.29	0.19	0.40	0.29	0.61	0.56	0.60	0.41	0.02	0.01	0.01	0.32	0.71	0.60	0.89	0.90	na	0.40
VA2	0.03	0.77	0.92	0.60	0.28	0.65	0.79	0.88	0.92	0.79	0.54	0.05	0.01	0.02	0.43	0.75	0.05	0.93	0.88	0.04	na

Table 2. Results of bivariate cointegration tests for delivered southern pine pulpwood logs, for the cointegrating relation specified as containing an intercept but no policy shock. Prices were undeflated prices and were transformed by the natural logarithm. Postal abbreviation codes were used to identify states and numbers (1, 2) to identify Timber Mart-South price regions within states.

X/Y	AL1	AL2	AR1	FL1	FL2	GA1	GA2	LA1	LA2	MS1	MS2	NC1	NC2	SC1	SC2	TN1	TN2	TX1	TX2	VA1	VA2
AL1	na	0.62	0.18	0.21	0.38	0.28	0.27	0.07	0.05	0.09	0.12	0.8	0.1	0.03	0.04	0.86	0.17	0.34	0.09	0.52	0.01
AL2	0.18	na	0.07	0.29	0.01	0.18	0.21	0.04	0.07	0.07	0.05	0.82	0.19	0.44	0.04	0.83	0.2	0.21	0.01	0.63	0.01
AR1	0.15	0.15	na	0.09	0.27	0.66	0.63	0.35	0.11	0.1	0.2	0.42	0.01	0.6	0.21	0.86	0.25	0.17	0.09	0.96	0.01
FL1	0.06	0.1	0.01	na	0.52	0.54	0.69	0.33	0.1	0.07	0.16	0.43	0.05	0.06	0.16	0.9	0.21	0.24	0.1	0.82	0
FL2	0.11	0	0.02	0.44	na	0.28	0.51	0.03	0.02	0.06	0.02	0.5	0.41	0.02	0.05	0.9	0.2	0.18	0	0.84	0
GA1	0.17	0.18	0.4	0.69	0.58	na	0.33	0.35	0.04	0.71	0.14	0.42	0.44	0.09	0.03	0.96	0.25	0.33	0.15	0.97	0
GA2	0.45	0.12	0.52	0.44	0.82	0.15	na	0.4	0.01	0.8	0.1	0.74	0.26	0.36	0.04	0.97	0.25	0.54	0.1	0.93	0
LA1	0.07	0.33	0.44	0.71	0.31	0.12	0.52	na	0.01	0.03	0.11	0.42	0.43	0.59	0.16	0.94	0.26	0.05	0	0.79	0.01
LA2	0.06	0.22	0.25	0.71	0.27	0.17	0.1	0.02	na	0.07	0.08	0.46	0.52	0.44	0.5	0.95	0.25	0.53	0.19	0.99	0
MS1	0.01	0.11	0.09	0.21	0.09	0.14	0.13	0.03	0.03	na	0.26	0.46	0.5	0.01	0.01	0.94	0.1	0.31	0.02	0.96	0
MS2	0.03	0.14	0.06	0.11	0.24	0.18	0.08	0.01	0.01	0.92	na	0.81	0.28	0.46	0.02	0.03	0.26	0.58	0.01	0.57	0
NC1	0.37	0.85	0.37	0.63	0.84	0.36	0.67	0.31	0.28	0.72	0.51	na	0.92	0.57	0.3	0.89	0.25	0.46	0.3	0.55	0.01
NC2	0.2	0.75	0.01	0.56	0.71	0.25	0.67	0.09	0.15	0.12	0.16	0.9	na	0.54	0.08	0.92	0.25	0.36	0.23	0.96	0
SC1	0.18	0.23	0.5	0.47	0.57	0.06	0.41	0.29	0.14	0.51	0.28	0.87	0.62	na	0.24	0.96	0.23	0.33	0.16	0.84	0
SC2	0.03	0.28	0.2	0.64	0.15	0.07	0.18	0.13	0.2	0.01	0.06	0.76	0.34	0.43	na	0.92	0.24	0.66	0.06	0.97	0
TN1	0.02	0.81	0.16	0.33	0.49	0.78	0.47	0.19	0.18	0.59	0.27	0.32	0.5	0.53	0.59	na	0.16	0.39	0.12	0.95	0.01
TN2	0.28	0.76	0.61	0.78	0.72	0.61	0.53	0.31	0.23	0.54	0.45	0.21	0.62	0.59	0.27	0.89	na	0.48	0.28	0.61	0.01
TX1	0.19	0.23	0.05	0.4	0.65	0.14	0.78	0	0.28	0.18	0.37	0.83	0.5	0.65	0.34	0.92	0.25	na	0.02	0.79	0.01
TX2	0.04	0.01	0.04	0.43	0.06	0.26	0.35	0	0.07	0.02	0.03	0.8	0.58	0.03	0.05	0.92	0.27	0.05	na	0.83	0.01
VA1	0.1	0.28	0.5	0.7	0.73	0.47	0.65	0.42	0.08	0.16	0.27	0.18	0.62	0.7	0.33	0.01	0.16	0.3	0.24	na	0.01
VA2	0.33	0.38	0.61	0.53	0.79	0.57	0.76	0.54	0.39	0.74	0.41	0.75	0.58	0.55	0.54	0.3	0.24	0.52	0.34	0.95	na

Table 3. Meta-analytic regression results under alternative cointegrating relation specifications for paired prices from delivered southern pine sawlog regions.

Model	Dependent Variable (P-value) Mean	St Dev	R^2	Intercept	$DIST_{yx}$	VOL_y	VOL_x	$CONC_y$	$CONC_x$
LINEAR									
Simple	0.367	0.331	0.203	0.056	0.262	−0.071	−0.113	−0.034	1.490
t-stat				1.033	4.359	−1.585	−2.768	−0.165	7.128
Policy	0.655	0.198	0.012	0.661	0.044	−0.023	0.013	−0.250	0.198
t-stat				15.887	1.093	−0.761	−0.451	−1.524	1.276
Policy PPI	0.392	0.301	0.159	0.163	0.049	0.054	−0.137	0.124	1.406
t-stat				2.786	0.861	1.251	−3.208	0.631	6.297
PPI	0.400	0.304	0.180	0.058	0.054	0.076	−0.056	0.353	1.723
t-stat				1.049	0.953	1.813	−1.357	1.799	8.261
LOGGED									
Simple	0.297	0.304	0.267	−0.015	0.390	0.040	−0.119	−0.469	0.835
t-stat				−0.381	8.639	1.186	−4.029	−3.418	4.705
Policy	0.681	0.161	0.126	0.556	0.096	0.144	0.047	−0.107	−0.271
t-stat				14.316	2.568	5.315	1.630	−0.684	−1.821
Policy PPI	0.592	0.172	0.104	0.468	0.127	0.108	0.019	−0.295	0.142
t-stat				12.172	3.392	3.394	0.675	−1.920	0.965
PPI	0.534	0.213	0.108	0.322	0.057	0.208	0.073	−0.061	0.206
t-stat				6.684	1.224	6.102	2.011	−0.326	1.124

distance, the greater the p-value (and hence the lower the likelihood of cointegration). Volume in the away market was positively related to the probability of cointegration, just as in the manner hypothesized by Faminow and Benson (1990), supporting the concept of a basing-point pricing system. Here, it is more likely that a series y is cointegrated with a series from region x if region x is a large-producing region. Results on market concentration indicate that the greater the amount of industrial production concentrated in a few firms in region x, the greater the likelihood that the price from region y is cointegrated with it. Alternatively, the smaller the amount of concentration in region y, the greater the likelihood that other series are cointegrated with it.

We select the best cointegration specification on the basis of overall model significance and the importance and significance of individual explanatory variables. In particular, we require that the most important variable (distance) be positive and significant. In linear cointegration specifications, this is true only for the simple model. In the logged specification, all but the PPI deflated version exhibit a positive and significant distance variable. We also consider the significance of concentration and volume variables. Here, the simple logged specification is clearly superior to other logged models. The simple linear model compares favorably. Not surprisingly, the simple linear and logged models also have the highest R^2 values.

Having determined the preferred specification, we return to the question of market identification. The relevant data set now consists of 420 p-values and their associated dependent and independent price series. Identifying cointegration can consist of specifying the correct cointegrating relation and a probability threshold. However, in this case, economic theory provides little guidance into which market corresponds to y (dependent) and which market corresponds to x (independent) in a particular cointegration test. For this reason, we evaluate both specifications. However, alternative specifications often provide contradictory results. We give each specification equal weight in the market determination process. Thus, the p-value employed in market determination for a pair of markets is the average of the p-values from two specifications. After transforming p-values in this manner, we have 210 market pairs. We limit further investigation to those pairs with average p-values less than 0.10. Limiting the data in this manner leaves 61 pairs of linked markets.

Visual inspection of the 61 remaining pairs leads to identification of 3 major and 2 minor overlapping regions in the southern pine sawlog market. The major regions are West, Central Gulf, and Southeast Atlantic. The West region includes Texas, Arkansas, Louisiana, and sometimes Mississippi and Eastern Tennessee. The Southeast Atlantic region includes Florida, South Carolina, and Georgia. Eastern Tennessee, North Carolina, and Virginia appear to be somewhat linked to the Southeast Atlantic region. However, it may also be correct to consider North Carolina and Virginia as a small, separate region (Northeast) that exhibits substantial overlap with the Southeast region. Overlap is consistent with the significant parameter estimate on the distance variable, which suggests that nearby prices are more likely to be cointegrated, other factors held constant. Finally, there is some evidence of a Southern Appalachian region. This region consists of Tennessee (region 1), South Carolina (region 1), North Carolina (region 1), Virginia (region 1), and Alabama (region 1). See Figure 1 for a diagrammatic representation of these regions.

Delivered southern pine pulpwood: After using chi-squared regressions to evaluate sawlog price cointegration specifications, we evaluate p-values arising from pulpwood cointegration tests in the same manner. The results of these tests are contained in Table 4. Note that distance is not statistically related to the probability that the null of no cointegration could be rejected in all cases except one; in that case, where cointegration was tested with a cointegration relation specification that deflated prices and included a Pacific Northwest policy shock dummy variable, the direction of influence was counterintuitive. Delivered pulpwood logs are lower in relative worth when compared to sawlogs; thus, the broad insignifi-

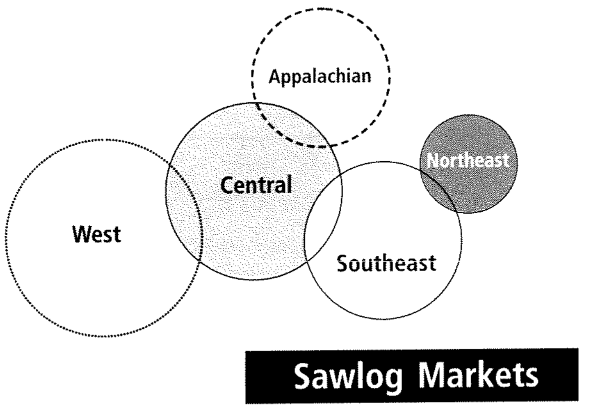

Fig. 1. Sawtimber Markets Diagram.

cance of distance across pulpwood model specifications is not surprising. It is likely that the high relative transportation costs associated with pulpwood lessen the influence of distance on pulpwood market integration. We speculate that any distance as great as the typical distance between the centroids of two adjacent regions would provide as much of a barrier to profitable product transport as a much larger distance.

Another notable factor from these regressions is the importance of "home" (the price in the dependent position of the cointegrating relation specification). It seems that concentration in the "home" market has a powerful influence on cointegration probability: the greater the concentration of output by a few firms in region y, the greater the likelihood that a comparison region's price would be cointegrated with it, providing some supporting for the market structure indicated by Murray (1995). Other notable factors from these regressions are the high explanatory power when modeling the simple linear and logged cointegration specifications.

Because cointegration in pulpwood markets is much less dependent on distance than it is in sawlog markets, we conjecture that the price movements within each series are more closely tied to the nonstationary demand and supply shocks experienced in local markets than are the corresponding sawlog series. Additionally, given the broad, regional integration in pulpwood, it might be concluded that market pulpwood prices in the South exhibit a closer relationship to broader demand variables. This kind of integration might be caused by the relative concentration of the pulp and paper industry. A few large firms may be able to coordinate price movements and shift production around the South within the firm to adjust for local supply shocks.

Here again we select the simple logarithmic specification as the preferred version of the cointegrating regression. As with sawlogs, we define integrated mar-

Table 4. Meta-analytic regression results under alternative cointegrating relation specifications for paired prices from delivered southern pine pulpwood log regions.

Model	Dependent Variable (P-value) Mean	Dependent Variable (P-value) St Dev	R^2	Intercept	$DIST_{yx}$	VOL_y	VOL_x	$CONC_y$	$CONC_x$
LINEAR									
Simple	0.375	0.305	0.267	0.592	−0.028	0.671	−1.765	−55.478	1.194
t-stat				12.486	−0.513	1.164	−2.977	−11.666	0.249
Policy	0.598	0.201	0.080	0.615	−0.001	−1.001	1.620	−6.010	−2.405
t-stat				19.882	−0.186	−2.428	4.061	−1.624	−0.675
Policy PPI	0.455	0.267	0.060	0.486	0.057	0.687	0.078	−26.292	−1.487
t-stat				10.705	0.959	1.098	0.134	−4.595	−0.281
PPI	0.453	0.266	0.060	0.543	−0.012	0.997	−0.497	−27.130	−1.668
t-stat				12.230	−0.203	1.639	−0.863	−4.876	−0.321
LOGGED									
Simple	0.342	0.285	0.192	0.433	0.033	0.632	−1.617	−40.015	1.528
t-stat				10.159	0.687	1.226	−3.087	−9.259	0.342
Policy	0.603	0.181	0.111	0.603	−0.004	−1.161	1.908	0.465	−5.699
t-stat				23.058	−0.133	−3.473	5.739	0.152	−1.895
Policy PPI	0.535	0.172	0.034	0.565	−0.0782	0.2201	1.027	−2.064	−5.004
t-stat				21.181	−2.249	0.647	2.982	−0.661	−1.622
PPI	0.561	0.220	0.082	0.467	0.053	1.424	2.125	−2.824	−9.838
t-stat				13.983	1.218	3.369	4.952	−0.716	−2.544

kets as those having average p-values <0.10. Limiting the data in this manner leaves us with 35 pairs of linked markets. Visual inspection of the remaining pairs seems to show that the southern pulpwood market is very poorly integrated from the perspective of local supply and demand shock transmission, but that the LOP may reasonably be expected to hold in three subregions. One region comprises an arc across the Gulf of Mexico coast and along much of the Atlantic coast. This could arise from the threat of arbitrage by rail or sea shipment to coastal pulp mills. The price series from Texas [2], Mississippi, Louisiana, and South Carolina [2] form a group that is consistently cointegrated with others in this region. It is likely that these regions have been less affected by internal shocks than other regions. States in the north and northeastern portions of the South are not generally related to each other or any of the major regions described. These states include North Carolina, Virginia, and Tennessee. See Figure 2 for a diagrammatic representation of these regions. The implication here is that the pulpwood log market is highly fragmented in its northern reaches and less so in its coastal areas. Coastal

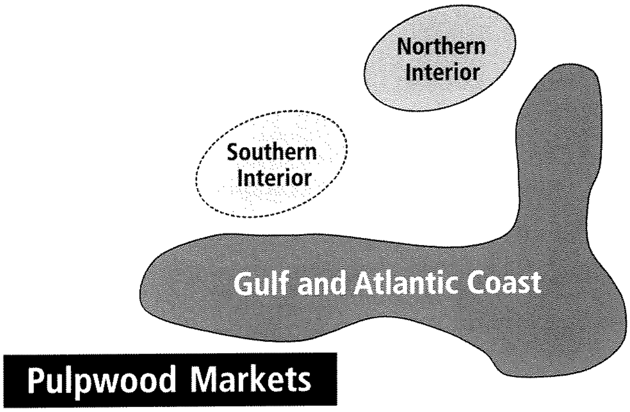

Fig. 2. Pulpwood Markets Diagram.

area prices may contain long-run relationships that are maintained not through arbitrage but through integrated output markets (e. g., pulp mills), whose various producing zones do share and transmit market shocks; in the north, even output markets are not integrated enough to hold long-run pulpwood prices together.

Conclusions

Cointegration tests provide a means for analysis of both the LOP and market integration. However, the results of such tests are heavily dependent upon correct specification of the cointegrating equation. In addition, cointegration tests can identify price linkages. However, these linkages may arise for a variety of reasons with vastly different implications. For example, price linkages arising from spatial arbitrage imply some degree of efficiency, whereas linkages arising from price collusion or only because output markets are well integrated imply quite the opposite of efficiency.

In this research, we have attempted to identify the structure of Southern softwood markets. In so doing, we have learned that the best cointegrating relation specifications for these markets are ones in which prices are nominal, transformed by the natural logarithm, and include no index of structural change linked to Pacific Northwest timber set-asides. Additionally, we have identified the influence that economic variables have on market integration. Meta-analysis indicates that distance is related, in a manner consistent with theory, to the finding of cointegration for sawlog markets, but it is not for pulpwood. Hence, at least some of the similar movement in sawlog prices can be linked to spatial arbitrage or its threat, while the same cannot be said for pulpwood. Further, these regressions show that relative volumes of production and market concentration significantly explain cointegration for sawlog, but not for pulpwood. Finally, the degree of market concentration matters especially for pulpwood, but its direction of influence on findings in favor of cointegration and the LOP is somewhat ambiguous.

Given these results, it is clear that markets and factors influencing markets are quite different for sawlogs and pulpwood. Perhaps consistent with this difference is that the spatial structure of these two commodity markets is different as well. These findings carry with them additional implications and therefore suggest avenues for further research.

First, given that sawtimber and pulpwood may be supply complements in the short run (pulpwood logs derive from sawtimber harvests) and substitutes in the long run (landowners can harvest either pulpwood on short rotations or sawtimber and pulpwood on longer rotations), the statistical relationships between them are likely to be complicated. Statistical studies that clarify understanding of their co-relationships, however, could yield major benefits to southern pine growers in the South, who need to predict long-run price trends of both products in order to make optimal investment and harvest decisions.

Second, our results on the kind of spatial arrangement of southern pine sawlog and pulpwood markets suggest that more precise timber inventory and market projection models (e. g., SRTS) may be possible, and these models would do a better job at evaluating the effects of local market shocks. Our results imply that local supply and demand shocks in sawlog markets are transmitted within the boundaries of our delineated "sub-markets" but not beyond those boundaries. Given that, natural catastrophes with concentrated regions of damage should affect prices throughout sub-markets but not beyond their boundaries. In contrast, shocks to local pulpwood supplies should not be spatially transmitted because pulpwood prices are more influenced by demand factors that shift in a coordinated fashion across the whole of the South, keeping prices together in the long run. However, shocks to pulp and paper markets may filter down quite easily across all coastal regions of the South but have uneven effects in other parts. Specific case studies, including more tests of market shock imprints like the one conducted by Prestemon and Holmes (2000), could include more direct tests of the existence of the sawlog and pulpwood log sub-markets delineated here.

received 14 May 2001 · accepted 21 March 2003

References

Abt RC, Cubbage FW and Pacheco G (2000) Southern forest resource assessment using the subregional timber supply (SRTS) model. Forest Products Journal 50 (4): 25–33.

Adams DM and Haynes RW (1980) The 1980 softwood timber assessment market model: structure, projections and policy simulations. Forest Science Monograph No. 22.

Alavalapati JRR Adamowicz WL and Luckert MK (1997) A cointegration analysis of Canadian wood pulp prices. American Journal of Agricultural Economics 79 (3): 975–986.

Ardeni PG (1989) Does the law of one price really hold for commodity prices? American Journal of Agricultural Economics 71 (3): 661–669.

Baffes J (1991) Some further evidence on the law of one price: The law of one price still holds. American Journal of Agricultural Economics 73 (4): 1264–1273.

Barrett CB (1996) Market analysis methods: Are our enriched toolkits well-suited to enlivened markets? American Journal of Agricultural Economics 78 (3): 825–289.

Baulch B (1997) Transfer costs, spatial arbitrage, and testing for food market integration. American Journal of Agricultural Economics 79 (2): 477–487.

Buccola ST (1985) Pricing efficiency in agricultural markets: issues, methods, and results. Western Journal of Agricultural Economics 14 (1): 111–121.

Buongiorno J and Uusivuori J (1992) The law of one price in the trade of forest products: Co-integration tests for U. S. exports of pulp and paper. Forest Science 38 (3): 539–553.

Dickey DA and Fuller WA (1979) Distribution of the estimators for autoregressive time series with a unit root. The Journal of The American Statistical Association 74 (366): 427–431.

Dykstra DP and Kallio M (1987) Introduction to the IIASA forest sector model. In: The Global Forest Sector: An Analytical Perspective (eds. M Kallio, DP Dykstra, and CS Binkley): 139–167. Wiley and Sons, New York.

Engle RF and Granger CWJ (1987) Co-integration and error correction: representation, estimation, and testing. Econometrica 55 (2): 251–276.

Espey M (1998) Gasoline demand revisited: an international meta-analysis of elasticities. Energy Economics 20 (3): 273–295.

Faminow MD and Benson BL (1990) Integration of spatial markets. American Journal of Agricultural Economics 72 (1): 49–62.

Gonzalo J (1994) Five Alternative methods of estimating long-run equilibrium relationships. Journal of Econometrics 60 (1/2): 203–233.

Goodwin BK and Schroeder TC (1991) Cointegration tests and spatial linkages in regional cattle markets. American Journal of Agricultural Economics 73 (2): 452–464.

Guan H and Munn IA (2000) Harvest restrictions: An analysis of new capital expenditures in the Pacific Northwest and the South. Journal of Forestry 98 (4): 11–17.

Haight RG and Holmes TP (1991) Stochastic price models and optimal tree cutting: Results for loblolly pine. Natural Resource Modeling 5 (4): 423–443.

Hall A (1994) Testing for a unit root in time series with pretest data-based model selection. Journal of Business Economics and Statistics 12 (4): 461–470.

Hänninen R Toppinen A and Ruuska P (1997) Testing arbitrage in newsprint imports to United Kingdom and Germany. Canadian Journal of Forest Research 27 (12): 1946–1952.

Hedges LV and Olkin I (1985) Statistical Methods For Meta-Analysis. Academic Press, Boston.

Holmes TP (1991) Price and welfare effects of catastrophic forest damage from southern pine beetle epidemics. Forest Science 37 (2): 500–516.

Howell M and Levins R (1998) Arkansas's timber industry: An assessment of timber product output and use, 1996. USDA Forest Service Research Bulletin SRS-28. Asheville, North Carolina.

Johansen S (1991) Estimation and hypothesis testing of cointegration vectors in Gaussian vector autoregressive models. Econometrica 59 (6): 1551–1580.

Johansen S and Juselius K (1990) The maximum likelihood estimation and inference on cointegration – with application to demand for money. Oxford Bulletin of Economics and Statistics 52 (2): 169–210.

Johnson TG Gober JR and Nix JS (1997a) Alabama's timber industry: An assessment of timber product output and use, 1995. USDA Forest Service Research Bulletin SRS-27. Asheville, North Carolina.

Johnson TG Jenkins A and Brown DR (1997b) North Carolina's timber industry: An assessment of timber product output and use, 1995. USDA Forest Service Research Bulletin SRS-18. Asheville, North Carolina.

Johnson TG Jenkins A and Haxby TS (1997c) Florida's timber industry: An assessment of timber product output and use, 1995. USDA Forest Service Research Bulletin SRS-13. Asheville, North Carolina.

Johnson TG Jenkins A Scrivani JA and Foreman JM (1997d) Virginia's timber industry: An assessment of timber product output and use, 1995. USDA Forest Service Research Bulletin SRS-19. Asheville, North Carolina.

Johnson TG Jenkins A Scrivani JA and Foreman JM (1997e) Georgia's timber industry: An assessment of timber product output and use, 1995. USDA Forest Service Research Bulletin SRS-14. Asheville, North Carolina.

Johnson TG Jenkins Stratton DP and Bischoff PS (1997f) South Carolina's timber industry: An assessment of timber product output and use, 1995. USDA Forest Service Research Bulletin SRS-16. Asheville, North Carolina.

Jung C and Doroodian K (1994) The law of one price for U. S. software lumber: A multivariate cointegration test. Forest Science 40 (4): 595–600.

Lang MG and Rosa F (1981) Price variation in direct and terminal markets for cattle and hogs: An Illinois case. American Journal of Agricultural Economics 63 (4): 704–707.

MacKinnon JG (1991) Critical Values for Cointegration Tests. In Long Run Economic Relationships (eds. RF Engle and CWJ Granger). Oxford University Press.

McNew K and Fackler PL (1997) Testing market equilibrium: Is cointegration informative? The Journal of Agricultural and Resource Economics 22 (2): 191–207.

Mulligan GF and Fik TJ (1989) Price variation in spatial oligopolies. Geographic Analysis 21 (1): 32–46.

Murray BC (1995) Oligopsony, vertical integration, and output substitution: welfare effects in US pulpwood markets. Land Economics 71 (2): 193–206.

Murray BC and Wear DN (1998) Federal timber restrictions and interregional arbitrage in U. S. lumber. Land Economics 74 (1): 76–91.

Nagubadi V Munn IA and Tahai A (2001) Integration of hardwood stumpage markets in the south-central United States. Journal of Forest Economics 7 (1): 69–98.

Norris Foundation (1977–1997) Timber Mart-South. The Daniel B Warnell School of Forest Resources, University of Georgia. Athens, Georgia.

Pantula SG Gonzalez-Farias G and Fuller WA (1994) A comparison of unit root criteria. Journal of Business and Economic Statistics 12 (4): 449–459.

Phillips PC (1998) New tools for understanding spurious regressions. Econometrica 66 (6): 1299–1325.

Prestemon JP and Holmes TP (2000) Timber price dynamics following a natural catastrophe. American Journal of Agricultural Economics 82 (1): 145–160.

Prestemon JP and Pye JM (2000) A technique for merging areas in Timber Mart-South data. Southern Journal of Applied Forestry 24 (4): 219–229.

Ravallion M (1986) Testing market integration. American Journal of Agricultural Economics 68 (1): 102–109.

Samuelson P (1952) Spatial price equilibrium and linear programming. American Economic Review 42 (3): 283–303.

Said SE and Dickey DA (1984) Testing for unit roots in autoregressive moving average models of unknown order. Biometrika 71 (3): 599–607.

Schnute J (1987) Data uncertainty, model ambiguity, and model identification. Natural Resource Modeling 2 (2): 159–212.

Smith VK and Kaoru Y (1990) Signals or noise? Explaining the variation in recreation benefit estimates. American Journal of Agricultural Economics 72 (2): 419–433.

Söderlind P and Vredin A (1996) Applied cointegration analysis in the mirror of macroeconomic theory. Journal of Applied Econometrics 11 (4): 363–381.

Stanley TD (1998) New wine in old bottles: A meta-analysis of Ricardian equivalence. Southern Economic Journal 64 (3): 713–727.

Stratton DP and Wright RC (1998) Tennessee's timber industry: An assessment of timber product output and use, 1995. USDA Forest Service Research Bulletin SRS-31. Asheville, North Carolina.

Stratton DP Howell DP and Wright RC (1998) Mississippi's timber industry: An assessment of timber product output and use, 1995. USDA Forest Service Research Bulletin SRS-29. Asheville, North Carolina.

Takayama T and Judge GG (1964) Equilibrium among spatially separated markets: A reformulation. Econometrica 32 (4): 510–523.
Tomek WG (1980) Price behavior on a declining terminal market. American Journal of Agricultural Economics 62 (3): 434–444.
Toppinen A and Toivonen R (1998) Roundwood market integration in Finland: A multivariate cointegration analysis. Journal of Forest Economics 4 (3): 241–266.
Uri ND and Boyd R (1990) Considerations on modeling the market for softwood lumber in the United States. Forest Science 36 (3): 680–692.